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B A S I S

Broadening Access and Strengthening Input Market Systems

Land market liberalization and the access to land by the rural poor: Panel data evidence of the impact of the Mexican Ejido reform

by

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1. Introduction

In the context of macro-economic liberalization, many developing countries have, during the last decade, made major strides towards replacing protectionist and interventionist policy regimes with greater openness, competition, and market orientation. Land markets seem, for various reasons, have been an exception to this rule. Even though the intensity of enforcement may have decreased, restrictions on the functioning and operation of land rental and sales markets such as prohibitions of share cropping, rent ceilings, regulations that confer different degrees of property rights on tenants, and the threat of expropriation unless land is put to proper use, continue to be on the books of developing countries all over the world.

While it is widely agreed that many of these restrictions (especially those on rental markets) may be associated with efficiency losses, failure to eliminate them is often justified by referring to the supposedly deleterious consequences on the poor of greater privatization of property rights and "unfettered" land market operation. However, very little empirical evidence exists to back such a statement and the policy debate is characterized by generalizations (e.g. failure to distinguish between land sales and rental markets) that tend to obfuscate rather than clarify. More conclusive research on this issue would be clearly desirable.

In fact, given the importance of the issue in a number of countries, and the intensity of the debate on the impact of increasing transferability of land in many developing countries, one is struck by the limited attention devoted to land market liberalization in the literature. The paucity of true land market liberalizations, and the data needed to make inferences on the impact of such policy changes, may explain the paucity of empirical studies on the subject. Still, compared, for example, to the voluminous literature on share tenancy, there is also little conceptual work that identifies the channels through which land market liberalization may affect the poor, thus offering not only a justification for empirical studies, but also some guidance on how to go about reducing restrictions on the functioning of land markets.

Such work would be particularly justified since restrictions which may have been essential to protect the poor in an environment where agricultural protection and subsidized credit propped up an economically irrational demand for land by large landlords may, in a liberalized environment, convert themselves into obstacles that prevent land access by the poor and thus greater efficiency as well as equity. However, this also implies that it is neither feasible nor desirable to focus only on land and neglect other factor markets. In particular, aiming to disentangle the interaction between land and credit markets is important not only to measure the impact of land market liberalization but also to provide on guidelines on how to sequence steps so as to ensure that measures to improve the functioning of land markets do not run counter to poverty reduction goals.

This paper aims to make an initial step in this direction by studying the impact of land rental market liberalization in the context of the Mexican *ejido* reforms which eliminated restrictions on land transactions that were imposed in the context of the various land reform episodes following the 1910 revolution. These rules, summarized in the Agrarian Code, stipulated that *ejido* land could not be sold, rented, or mortgaged, that usufruct rights would be contingent on occupation and cultivation, and that subdivision, even in the context of inheritance, would be prohibited. Intergenerational succession and increased off-farm employment had rendered the legal framework increasingly dysfunctional, evidenced by widespread neglect and circumvention. Nonetheless, the fact that such rules continued to affect the majority of Mexico's rural population, made the issue of their formal abolition into a highly controversial policy issue.

The paper is structured as follows. Based on brief description of the background, character, and implementation modalities of Mexican reforms, section 2 elaborates a conceptual model to generate empirically testable hypotheses regarding the impact of land market liberalization on land demand and the derivative of cultivated land with respect to owned land of credit constrained and unconstrained households in different regimes (renting in and renting out) depending on their factor endowment. We note that, in a world without capital constraints, land market liberalization would unambiguously increase demand by land-poor households as well as supply of land by large land owners, thus improving overall productivity and household welfare. With binding credit constraints and unchanged credit access, one would still expect a positive impact of land market access, although the magnitude would be reduced as compared to the unconstrained case. If, as a result of reforms, access to credit changes as well, the net impact will depend on this overall change and the relationship between access to liquidity and households' land endowment.

Section 3 uses descriptive statistics from 1990, 1994, and 1997 to show that, in the case of Mexico, liberalization of land markets and tightening in credit access occurred simultaneously. They suggest that reforms increased small farmers' access to land but at the same time also greatly reduced their access to credit. This provides a justification for more detailed examination using panel data from 1994 and 1997 for about 1300 households.

To this end, and to be able to accommodate the non-linear character of the relationship between land demand and land endowment, we apply semi-parametric methods in section 4. For the credit unconstrained case, estimation of the non-parametric part of the land demand function in endowment space points towards a shift that is consistent with the predictions of the theoretical model – in the second period access to land by the land-poor and supply by the land-rich have increased between 1994 and 1997. To empirically control for credit access, we use a conditional maximum likelihood approach, as suggested by Kyriazidou (1997). In addition to finding (from the conditional logit) that participation in informal organizations (which serves as an exclusion restriction) reduces the probability of households' being credit constrained, this allows us to obtain conditional non-parametric estimates of the relationship between land demand and land endowment. In addition to confirming the general conclusion from the unconditional model, these estimates are also statistically significant at the 90% confidence level.

Returning to the general question that motivated the paper, we find that, even though, in a world with binding capital constraints, land market liberalization could hurt the poor, this does not seem to have been the case in Mexico. To the contrary, we observe an upward shift in land demand by the poor and landless, together with a downward shift in land demand by those well-endowed with this factor of production. In the end, and even though many of the regulations appear have been disregarded earlier, land rental market liberalization seems to have created a win-win situation that allowed to increase both intensity of agricultural cultivation and the welfare of the rural poor. Contrary to widespread belief, well-sequenced episodes of land market liberalization appear to have potential to help, rather than hurt, the poor. Other countries where restrictions on the functioning of land rental markets continue to be enforced more stringently than they have been in Mexico might be able to significantly improve the welfare of the rural poor by eliminating restrictions on the operation of such markets.

2. Mexico's Ejido System and the 1992 reforms

In this section we provide a more detailed description of the policy changes and their expected impact which serves as the motivation for the paper. We then develop a micro-theoretic model in Section 3 to derive testable hypotheses regarding the impact of land and credit market liberalization on household decisions. Sections 2.1 and 2.2 describe background and present some preliminary descriptive statistics that motivate the conceptual work developed in Section 3.

4.1 Background

The pre-1992 model of Mexico's ejido system was a product of the land reform program instituted following the 1910 peasant revolution. Land ownership was granted to communities or groups of families, the *ejido*, who, in return, allocated permanent, inalienable and inheritable land use rights to group members, or *ejidatarios*.[1] Within the ejido, *parcelized* and *common* lands (*tierras parceladas* and *comunales*) were distinguished. Individual usufruct rights to maximum of 20 hectares of parcelized land was obtained through an indivisible title (*titulo parcelario*), contingent on direct cultivation. Common lands were used collectively, with details governed by the ejido assembly.

To appreciate the importance of the *ejido* sector, note that, the early 1990s, over 15 million peasants and their families (about 21% of Mexico's population) were registered as *ejidatarios* in approximately 29,000 ejidos and agrarian communities throughout the country (INDA, 1996). *Ejido* land accounted for almost 55 percent of Mexico's arable land (Jones and Ward, 1998) and 70 percent of its forest cover. Although the majority of *ejido* lands were held as commons (75%), almost 85% of ejidatarios had access to individual parcels, the average size of which was 9.5 hectares per household (Thompson and Wilson, 1994).

For reasons of efficiency as well as equity, the agrarian reform sought to allocate parcellized land to owner-cultivated family farms. To prevent re-aggregation of land or re-emergence of "feudal" production relations and absentee landlordism, the *Agrarian Code* made possession of parcelized land contingent on owner-cultivation and prohibited the hiring of wage labor as well as the rental, sale, or mortgaging of land. Though usufruct rights were inheritable,

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[1] Indigenous communities recognized as such were given a different status, under the rubric of agrarian communities.

subdivision upon inheritance was prohibited to prevent land fragmentation.[2] As a result, new groups emerged such as the *avecindados*, residents in ejido communities without access to ejido land, and *posesionarios*, who have legal usufruct over ejido land, but not the voting rights and privileges of an ejidatario.

This increasing differentiation of the countryside (Harvey, 1994), together with the emergence of migration and off-farm employment opportunities, greatly increased the scope for gains from land rental transactions. Indeed, it is estimated that between 50-90% of *ejido* lands in Mexico's irrigated northwest, and about 35 to 50% in the rain-fed areas were subject to some form of (illegal) rental arrangements (Yates 1981).[3] The substantial amount of land rentals notwithstanding, the literature is very clear on the fact that the threat of eviction for *ejido* members who rented out land continued to exist (Finkler 1978; Heath (1992). Eliminating this risk was the main purpose of the legal reforms introduced in 1992.

These reforms have been the subject of intense debate. Supporters claimed that, by inducing land rich farmers to increase their supply of land to the rental market, the reforms could increase the efficiency of resource use and, at the same time, improve land access by the rural poor. Critics argued that, without previously mitigating the effects of imperfections in credit, input and output markets that give large farmers a competitive advantage over small family farms, the reforms would result in a large-scale sell-out by the rural poor, perhaps inducing a new wave of land re-concentration. This paper intends to shed light on this debate by examining, both conceptually and empirically, these effects using panel data from 1994 and 1997. We argue that, to understand the impact of property rights reform (which, in a neoclassical framework of perfect markets, would have an unambiguously positive impact on productivity and household welfare) in a context of imperfect markets, it is critical to consider the simultaneous changes in the financial system. In the remainder of this section we explore the data descriptively to acquire a broad view of the changes experienced by ejidatatrios since the 1992 reforms in terms of land rental market participation and access to credit.

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[2] Non-ejido members were also prohibited from entering into contracts with the ejido on cultivation of common lands (de Janvry, et al, 1997).

[3] DeWalt and Rees (1994) review several studies which provide evidence of the extent of rental and sales prohibition violations.

4.2 *Data and descriptive analysis*

The data analyzed in this study are from three surveys of households in the Mexican ejido sector. The first survey was carried out by SARH-CEPAL in 1990, and is representative of the ejido sector at the national, state and rural development district levels. This was a nationwide survey, including 5,007 ejidos and 35,090 ejidatarios and community members. Two follow-up surveys, one in 1994 and one in 1997, were conducted in a sub-sample of 275 ejidos previously surveyed in 1990. In average, five households per ejido were surveyed in 1994 and again in 1997, resulting in a two-year panel of approximately 2600 observations (1300 households per year). The objective of the follow-up surveys was to identify the differentiated impact of the reforms at the ejido and ejidatario levels.

Table 1: Predominant types of land tenure in the Mexican ejido in 1990, 1994 and 1997

	1990[1]		1994[2]		1997[2]	
	% of ejidatarios	Average size (ha)	% of ejidatarios	Average size (ha)	% of ejidatarios	Average size (ha)
Ejidatarios with land:						
Owned	98.3	9.8	97.0	10.2	95.5	12.3
Not owned	4.7	8.4	8.5	5.1	9.9	10.3
Rented	1.2	5.5	3.0	5.4	4.2	11.1
In partnership	0.6	5.6	1.2	4.1	2.5	6.9
Loaned	1.0	4.6	2.3	5.5	3.8	8.9
Granted	1.2	14.7	0.1	4.0	0.0	.
Other	0.9	7.7	2.2	4.0	0.2	9.1
Rented to others	1.4	2.6	6.4	5.2	7.8	6.5

[1]Source: de Janvry, Gordillo and Sadoulet (1997). [2]Own calculation.

As seen in Table 1, the increasing percentage of *ejidatarios* engaging in land rental transactions (either using land owned by others or renting it to others) suggests that, between 1990 and 1997, land rental markets in the Mexican *ejido* sector became substantially more active. The fact that, according to tables 1 and 2, most of the growth in rental activity occurred between 1990 and 1994 suggests that at least part of the "rental boom" may be explained by the fact that, with rentals now being legal, farmers who had engaged in illegal rental transactions earlier admitted to doing so in the 1994 survey. Nonetheless, the fact that growth persisted between 1994 and 1997, albeit at lower rates, suggests that the reforms may have had a structural impact.

Table 2: Land market by farm size and geographic region.

	1990[1]		1994[2]		1997[2]	
	% of Ejidatarios with land					
	Not owned	Rented to others	Not owned	Rented to others	Not owned	Rented to others
ALL	4.7	1.4	8.5	6.4	9.9	7.8
<2	3.9	0.2	2.4	9.9	2.5	12.2
2-5	3.5	1.5	8.0	4.2	10.4	2.4
5-10	3.3	0.8	6.8	5.4	6.9	6.2
10-18	5.4	1.5	12.8	6.6	12.2	7.2
>=18	15.4	4.3	25.8	7.8	23.4	4.2
<5	3.7	0.8	5.7	6.6	7.5	9.8
>=5	6.0	1.6	12.2	6.2	12.1	6.0

[1]Source: de Janvry, Gordillo and Sadoulet (1997). [2]Own calculation.

Also, as shown in Table 2, while household participation in rental markets between 1990 and 1994 increased for both large and small farmers (with more or less than 5 ha NRE), between 1994 and 1997 the increase in participation was mostly observed in the small farm sector (from

5.7 to 7.5% using land not owned, and 6.6 to 9.8% renting to others). Therefore, these simple descriptive results suggest that small farmers did gain more access to land through the rental market since the 1992 reforms. In the econometric analysis that follows, we will investigate this hypothesis more thoroughly.

Given the importance of changes in credit access on households' participation in land rental markets, tables 3-5 present some descriptive results on changes in access to credit by farm size. They suggest that access to formal sources of credit dropped drastically from 1994 to 1997. The percentage of *ejidatario* households that used formal credit fell from 25 to 11 percent. Those who used credit shifted away from Pronasol (a government program that was dropped) to Banrural. Further, in 1994 large holders (> 5 has NRE) received credit with a higher frequency, primarily due to discrepancies in access to Banrural. In 1997, however, access to Banrural became more egalitarian, primarily due to a large increase among landholders in the 2-5 has category.[4] As a consequence access to credit overall in 1997 seems to have become more equal between large and small holders.

Although access to Banrural increased over the 1994-1997 period, loan amounts, in constant 1994 pesos, dropped notably. In general, the amount of money available for ejido agriculture from formal sources fell drastically over this period (Table 4). While in 1994 Banrural granted 48 pesos/ha in loans (over the whole sample), and formal sources overall 134 pesos/ha, in 1997 this dropped to 19 pesos/ha for Banrural, and 40 overall. Pronasol dropped from 47 to 1 peso/ha, and commercial credit from 21 to 4 pesos /ha. In terms of Banrural, the disbursement of credit became more egalitarian, though overall this was offset by the near disappearance of Pronasol, which favored small holders.

The drop in credit available to ejido household was due not only to more restricted access, but also lower real loan amounts. In Table 5, we present average loan sizes over those ejidatarios who received loans. The drop in loan amount is uniform across sources and farm sizes, with the exception of other sources. The average Banrural loan fell from 1135 pesos/ha in 1994 to 278 in 1997. Overall the drop was from 534 pesos/ha in 1994 to 377 in 1997.

In summary, while loans amounts appear to have been reduced for all farm sizes since the implementation of reforms, access to credit in the ejido sector appears to have become more egalitarian instead of more wealth biased as some would have expected. This may partially explain the increase in the percentage of small farmers (2-5 hectares NRE) renting land in, and the increase in the percentage of medium farmers (5-18 hectares NRE) renting land out. It perhaps also explains the decrease in rentals by the smallest group of farmers (<2 hectares NRE), for whom a further drop in access to credit may have led them to reduce their area of cultivation even further, despite possibly falling rental rates. Finally, larger farmers (>18 hectares NRE) who perhaps have other sources of liquidity and therefore were not affected by the credit crunch, appear to have taken advantage of falling rental rates to increase their demand for rented land. In Section 3 we investigate these changes more thoroughly by employing econometric methods.

[4] A large portion of this increase in access to Banrural took place in the Gulf region. It was most likely politically motivated due to nationally important state elections, and the election of a former Secretary of Agrarian Reform as Governor of another state (Davis, 1999).

Table 3. Percentage of ejidatarios who use formal credit, by farm size, 1994-1997

	1994									1997								
		farm size (ha NRE)									*farm size (ha NRE)*							
	overall	*0*	*e-2*	*2-5*	*5-10*	*10-18*	*>18*	*<5*	*>5*	*overall*	*0*	*e-2*	*2-5*	*5-10*	*10-18*	*>18*	*<5*	*>5*
Number of observations	1308	24	186	374	306	257	162	560	724	1308	24	186	374	306	257	162	560	724
Public sources																		
Banrural	4	0	1	1	6	8	6	1	7	7	0	3	9	7	5	11	7	7
Pronasol	18	7	17	18	20	14	22	17	18	1	0	1	2	1	0	2	1	1
Other government agencies	1	3	0	1	1	0	1	1	1	1	0	2	1	1	1	2	1	1
Formal private sources																		
Commercial banks	1	0	0	0	4	0	1	0	2	1	0	0	1	0	1	1	0	1
Other sources	1	1	0	0	2	3	1	0	2	1	0	0	0	1	2	4	0	2
Overall	25	16	18	21	32	26	30	20	29	11	0	5	13	9	9	18	10	11

Table 4. Average loan size over total sample, by farm size, 1994-1997. (1994 Pesos)

	1994									1997								
		farm size (ha NRE)									*farm size (ha NRE)*							
	overall	*0*	*e-2*	*2-5*	*5-10*	*10-18*	*>18*	*<5*	*>5*	*overall*	*0*	*e-2*	*2-5*	*5-10*	*10-18*	*>18*	*<5*	*>5*
Number of observations	1308	24	186	374	306	257	162	560	724	1308	24	186	374	306	257	162	560	724
Public sources																		
Banrural	48	0	11	17	66	77	97	15	76	19	0	8	18	32	23	8	14	24
Pronasol	47	22	80	70	35	22	21	74	27	1	0	2	1	2	0	0	2	1
Other government agencies	4	1	0	10	3	0	0	7	2	3	0	7	4	0	1	5	5	1
Formal private sources																		
Commercial banks	21	0	0	5	77	5	9	3	36	4	0	0	11	0	3	2	8	2
Other sources	14	84	2	5	8	38	11	4	19	12	0	0	2	29	23	1	1	21
Overall	134	108	93	107	187	141	139	102	160	40	0	17	37	64	50	17	30	49

3. Modeling the impact of policy reforms on household land allocation decisions

In this section we develop a micro-theoretic model to derive testable hypotheses regarding the impact of land and credit market liberalization on household decisions. Our main goal is to show that, when all else remains constant, land market liberalization reforms will decrease the dependence of cultivated area A on land endowments T. In other words, reforms would lead to an increased separability between land demand A and land endowments T.

Section 3.1 describes the model setup. In sections 3.2 we derive the expected impacts of reforms on households that may operate in different participation regimes in the land, labor and credit markets. These expected impacts are the hypotheses to be empirically tested in Section 4.

3.1 The basic production model

To derive testable hypotheses, we develop a stylized model to motivate the empirical approach employed in Section 4. The key element in the model is that prior to the implementation of reforms, it is assumed that households renting land out faced a probability ρ of having their land confiscated by the ejido authority. To model the effect of confiscation threat, we assume that households maximize the sum of current consumption c and the expected value of a twice-differentiable quasi-concave utility function of terminal wealth, denoted $V(M)$. In $V(M)$, M denotes terminal wealth.[5] By assumption, before reforms M is a random variable for households renting land out, since they face the risk of having their land ownership rights revoked.

We assume further that ρ is a decreasing function of the ratio of cultivated to owned land (A/T). That is, the higher the share of a household's total land endowment rented out, the more visible the transaction is to others, and the higher is the probability of confiscation. Thus, we assume that ρ is given by $\rho=\rho(A/T)$, where $\rho(A/T)$ is differentiable everywhere, except at the point where $A=T$. Moreover, we postulate that $0<\rho(A/T)<1$ for $A<T$, but $\rho(A/T)=0$ when $A>T$. Moreover, we assume that $\rho'<0$, $\rho''>0$ for $A<T$, and that $\rho'=\rho''=0$ for $A>T$.

[5] The results would not change if we assumed, instead, that households maximize the sum of the utility of current consumption $U(c)$, and the expected utility of ending wealth.

Therefore, the probability of getting caught renting land out increases with *A/T*, at a decreasing rate, and the expected utility of ending wealth is given by:

$$W(A,T;M)=\begin{cases}\rho\left(\frac{A}{T}\right)V(M-T)+\left[1-\rho\left(\frac{A}{T}\right)\right]V(M), & for \quad A<T\\ V(M) & otherwise\end{cases}$$

where T is the household's land endowment. The household's objective function is therefore given by:

$$c + W(A,T;M).$$

Farmers face a constant-returns to scale production technology, represented by a twice differentiable and convex production function $Y=F(E;A)$, where A is the area cultivated, and E is the amount of labor effort employed into A. Because labor contracts are affected by moral hazard, the level of effort supplied by hired labor depends on the intensity of supervision by household workers (Eswaran and Kotwal, 1985). Therefore, we assume that the level of effort is given by the following function:

$$E \equiv \left(X_f + X_h\right)\left[\frac{X_f}{X_f + X_h}\right]^{1-\gamma}, \tag{1}$$

where X_f and X_h are the amounts of family and hired labor employed, and $\gamma\in[0,1]$, is an exogenous labor-extraction parameter that determines how effectively a farmer is able to extract effort from hire labor. If $\gamma=1$, the farmer is very effective, and hence, hired and family labor are perfect substitutes. On the other hand, if $\gamma=0$, the household is not able to extract any effort from hired labor. Note that (1) implies that when only family labor is employed, the level of effort is identical to the amount of family labor (i.e., $E=X_f$).[6]

[6] Note that because the effort function E(.) is homogeneous of degree one in Xf and Xh, and F(.) is homogeneous of degree one in E and A (because of CRTS), F(.) will be homogeneous of degree one in Xf and Xh.

Farm households are endowed with $\overline{L}$ units of time that can be either sold to outside employers or employed in their own farms. We assume that the market wage rates paid to hired labor and received for off-farm work are the same and equal to *w*. Households willing to take the risk of land confiscation may rent out part or all of their land endowments to others. They can also rent land in from others without any risk of punishment by the ejido authority. The market rental rate *r* is assumed to be identical for both land rented in and out. All prices are normalized in terms of the price of the consumption good *c*. Output from farm production is sold in a competitive market.

In addition to wage or rental income, households may finance production by borrowing in a competitive credit market. We assume that, because of either interest rate ceilings or adverse selection problems, interest rates are the same for all farmers, and credit market may not clear due to rationing. This implies that at the observed market interest rate, each household *i* has access to a fixed amount of credit S_i. We start our analysis by assuming that S_i does not bind. Later we relax this assumption to explore the combined impacts of land and credit market liberalization.

To summarize, the farm household problem is to:

Maximize: $c + W(A,T;M)$

w.r.t: X_f, X_h, A, and X_o.

s.t.:

$$c = F(E, A) - rA - wX_h + rT + wX_o$$

$$\overline{L} \geq X_f + X_o$$

$$E = \left(X_f + X_h\right)\left[\frac{X_f}{X_f + X_h}\right]^{1-\gamma}$$

$$W(A,T;M) = \begin{cases} \overline{\rho}\omega\left(\frac{A}{T}\right)V(M-T) + \left[1 - \overline{\rho}\omega\left(\frac{A}{T}\right)\right]V(M), & \textit{for} \quad A < T \\ V(M) & \textit{otherwise} \end{cases}$$

$$rA + wX_h \leq S + wX_o + rT$$

$$X_f, X_h, A, X_o \geq 0 \qquad ,$$

In the appendix we derive the first order necessary conditions for an optimal solution to the programming problem above. These solutions are demand functions that will take different forms depending on the household's participation regime in the labor, land or credit markets. For each of these regimes, we are interested in predicting the impact of reforms on the relationship between A^*, the household's specific optimal demand for cultivation area, and T, the household's predetermined land endowment. That is, for each participation regime, we are interested in predicting the impact of reforms on the derivative $\frac{\partial A^*}{\partial T}$.

3.2 The impact of reforms in the absence of liquidity constraints

The figures below depict this effect of reforms for liquidity-unconstrained and -constrained households. Figure 1 depicts the impact of land market liberalization for households with unlimited access to liquidity at the market interest rate. That is, for households that are price-rationed and not quantity-rationed in the credit market. The optimal choice of cultivated area for households in this regime is denoted by $A^* = A^u_{j,A>T}\left(T; w, r, \overline{L}\right)$, for tenant-farmers, and $A^* = A^u_{j,A<T}\left(T; w, r, \overline{L}\right)$, for landlord farmers. The superscript u indicates that the household is

operating under a liquidity-unconstrained regime. The subscript j=0, 1, denotes pre- and post reform periods, respectively.

As seen in Figure 1, and proven in the appendix, $A^u_{j,A>T}\left(T;w,r,\overline{L}\right)$, does not depend on T, both before and after reforms. That is, $\frac{\partial A^u_{j,A>T}}{\partial T}=0$, for j= 0 or 1, which implies that the demand for area cultivated of liquidity-unconstrained households will be completely separable from land endowments, both before and after reforms. That is, reforms would have no effect on $\frac{\partial A^u_{j,A>T}}{\partial T}$ for liquidity-unconstrained tenant farmers.

On the other hand, while the area demanded by liquidity-unconstrained landlord-farmers, i.e., $A^u_{j,A<T}\left(T;w,r,\overline{L}\right)$, will not depend on T after reforms (i.e., $\frac{\partial A^u_{j,A<T}}{\partial T}=0$, for j=1), it should depend positively on T before reforms (i.e., $\frac{\partial A^u_{j,A<T}}{\partial T}>0$, for j=1). That is, while farmers in this regime would have liked to have rented out T-A_0, because of the threat of land confiscation, they choose to supply only $T-A^u_{j,A<T}$, such that $A^u_{j,A<T}>A_0$. Therefore, before reforms $A^u_{j,A<T}\left(T;w,r,\overline{L}\right)$ is not separable from T. It increases as T increases.

Thus, as shown in the appendix, by eliminating the threat of land confiscation, the 1992 reforms would have had the effect of increasing the separability between A^* and T. As a result, we would observe a downward rotation of $A^u_{j,A<T}\left(T;w,r,\overline{L}\right)$ which is shown in Figure 1.[7] That is:

(2) $$\frac{\partial A^u_{0,A>T}}{\partial T}>\frac{\partial A^u_{1,A>T}}{\partial T}=\frac{\partial A^u_{1,A<T}}{\partial T}=\frac{\partial A^u_{0,A<T}}{\partial T}=0\,.$$

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[7] Everything else being held constant, under the assumptions of the model, this rotation would have made

$$\frac{\partial A^u_{1,A>T}}{\partial T}=\frac{\partial A^u_{1,A<T}}{\partial T}0\,.$$

This rotation implies in an increase in the supply of land to the rental market by land-abundant households, which would therefore benefit the land-poor through lower rental rates. Holding everything else constant, this fall in rental rates would cause an upward shift on the "free-market" demand for land from A_0 to A_1, as shown in Figure 1.

In summary, by increasing the separability between land owned and land cultivated, rental market liberalization reforms should lead to a factor price equalization by shifting land from land-abundant/labor-poor households to labor-abundant/land-poor ones.

3.3 The impact of reforms for liquidity constrained households.

We denote the demand for area cultivated of liquidity-constrained households by $A^* = A^c_{j,A>T}(T;w,r,\overline{L},S)$, for tenant-farmers, and $A^* = A^c_{j,A<T}(T;w,r,\overline{L},S)$, for landlord-farmers, where the superscript *c* denotes the liquidity-constrained regime. Note that, in contrast to the liquidity-unconstrained case, A^* is now a function of the household-specific access to liquidity denoted by *S*.

Would the reforms have a similar (qualitative) impact on liquidity-constrained households? As shown in Figure 2, and proven in the appendix and, the answer appears to be yes. This assumes, of course, that access to liquidity (i.e., *S*) remains unchanged. Figure 2 summarizes the comparative static results presented in the appendix under this assumption of unchanged *S*. As shown, because tenant-farmers must use up some liquidity to pay for land rentals, the relationship between A^* and *T* will be positive for both tenant- and landlord-farmers, and both before and after reforms (i.e., $\frac{\partial A^c_{j,A>T}}{\partial T}, \frac{\partial A^c_{j,A<T}}{\partial T} > 0$, for j=0 and 1). However, as shown in Figure 2, before reforms we would observe a "kink" at $A^*=T$, since $\frac{\partial A^c_{0,A<T}}{\partial T} > \frac{\partial A^c_{0,A>T}}{\partial T}$. In the appendix we show that reforms would have a similar *ceteris paribus* effect of rotating $A^c_{j,A<T}(T;w,r,\overline{L},S)$, thereby eliminating this "kink". Thus, the impact of reforms on the relationship between A^* and *T* for liquidity-constrained households can be summarized by:

(2') $$\frac{\partial A^c_{0,A>T}}{\partial T} > \frac{\partial A^c_{1,A>T}}{\partial T} = \frac{\partial A^c_{1,A<T}}{\partial T} = \frac{\partial A^c_{0,A<T}}{\partial T} > 0\ .$$

In this paper we examine the evidence of increased separability between A^* and T by testing the changes in slopes implied by (2) and (2'). In the following section we discuss the empirical strategy employed to identify the net effect of lifting rental prohibitions on this separability.

4. Econometric Analysis

Expressions (2) and (2') derived in Section 3 summarize the main hypotheses to be empirically tested in this section. They essentially state that land market liberalization in Mexico will have two effects in the relationship between A and T: (i) It will increase the separability between A and T (by "flattening" the *ceteris paribus* relationship between A and T); (ii) It will decrease the difference between the slope $\frac{\partial A^*}{\partial T}$ for households renting land in and households renting land out. These effects are predicted for both liquidity-constrained and unconstrained households.

We start testing these hypotheses by performing a semi-parametric regression of A on T that controls for several household characteristics (e.g., labor endowment), but does not control for access to credit S. As discussed below, this is a valid test only under the strong assumption that the relationship between access to credit S and land endowments T did not change over time. Since, as indicated by the descriptive analysis in Section 2, the relationship between S and T is likely to have changed between 1994 and 1997, we perform further econometric tests utilizing a sub-sample of households identified as credit-constrained, for which we can control for access to credit S in both periods. That is, for this sub-sample of credit-constrained households, we regress A on T, controlling for S and other household characteristics. Samples selectivity is dealt with by employing an instrumental variable approach developed in Kyriazidou (1997).

4.1 Semi-parametrically testing the impact of reforms (not controlling for access to credit)

As indicated before, our aim is to infer on the impact of reforms on changes in the relationship between A and T over time. Thus, we specify the following empirical model to represent this relationship:

$$A_{it} = g_t(T_{it}) + \beta S_{it} + \gamma' \mathbf{Z}_{it} + e_{it} \tag{3}$$

where $g_t(.)$ is a conditional mean function that may change with time, S_{it} is the amount of liquidity available to household i in period t, $\mathbf{Z}_{it}$ is a vector of household demographic characteristics (household labor force, age and gender of household's head, maximum education in the household, etc.), and e_{it} is a random disturbance such that $E[e_{it}|T_{it},S_{it},Z_{it}]=0$. While our ultimate inference interest lies on examining changes in the conditional mean function $g_t(.)$ between 1994 and 1997, we start with an unconditional analysis that also lets S_{it} depend on T_{it}. Hence, we specify

$$S_{it} = h_t(T_{it}) + \delta' \mathbf{Z}_{it} + u_{it}, \tag{4}$$

where $h_t(.)$ is a conditional mean function describing the relationship between access to liquidity and land endowments T_{it}, and $E[u_{it}|T_{it},Z_{it}]=0$. As indicated by the t subscript, this relationship may have changed between 1994 and 1997. That is, access to liquidity may have become more or less dependent on land wealth.

After substituting (4) into (3) we obtain the reduced form:

$$A_{it} = m_t(T_{it}) + \alpha' \mathbf{Z}_{it} + \varepsilon_{it}, \tag{5}$$

where $m_t(T_{it}) \equiv g_t(T_{it}) + \beta h_t(T_{it})$, $\alpha \equiv \gamma + \beta\delta$, and $\varepsilon_{it} \equiv e_{it} + \beta u_{it}$. Thus, changes in the conditional mean function $m_t(T_{it})$ between 1994 and 1997 reflect changes in the expected relationship between land demand A and land endowments T, which can be caused by land market liberalization reforms and/or differential changes in access to liquidity S_{it} for households with different land endowments T. Therefore, a test for changes in the relationship between A and T performed by testing for changes in $m_t(T_{it})$ over time, will only be a valid test for the impact of liberalization on $g_t(T_{it})$ if $\partial h_t(T_{it})/\partial T_{it}=0$.

As indicated by the pictures in Section 3, the relationship between land demand and endowments is likely to be highly non-linear. Thus we opt to estimate equation (5) semi-parametrically where $m_t(T_{it})$ is estimated via kernel regression methods and the vector of parameters α is estimated via linear regression.[8]

Estimation strategy and results

The non-parametric component of (5) is estimated via a kernel regression using the univariate Epanechnikov kernel function. Bandwidths for the kernel function were selected via cross-validation techniques.[9] Panel A in Figure 3 below displays the estimated $m_t(T_{it})$ for both 1994 and 1997. The shapes of both curves resemble the shape of the pre-reform curve depicted in Section 3. As can be seen, households endowed with less than approximately 3 hectares tend to operate as tenants. Their expected land demand is above the 45^o line, which suggests that they need to rely upon the land rental market to supplement their low land endowments. In contrast, the expected demand curve for households endowed with more than 3 hectares lies below the 45^o line, which indicates that they tend to cultivate less than their endowments and therefore rent out some excess land.

A comparison between the estimated relationship between *A* and *T* for 1994 and 1997 would suggest that reforms did indeed have the combined effect of increasing the supply of land by large owners, and that this increased supply appears to have been partially captured by the landless and the land poor. This is suggested by the upward shift on the demand curve of households owning less than 5 hectares, and the downward shift for households endowed with more than 5 hectares. However, these shifts are not statistically significant at the 10% level, as the 90% confidence regions around both curves overlap everywhere.[10] Accordingly, from the unconditional analysis above, we cannot reject the hypothesis that the combined land and credit market reforms had no effect on the levels of land demand for either large or small farmers in the Mexican ejido sector.

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[8] See Blundell and Duncan (1998) for details on estimation methods that combine kernel and linear regression techniques.

[9] The bandwidths used for 1994 and 1997 were h_{94}=5.2 and h_{97}=6.8, respectivelly.

Table 6 presents the estimation results for the parametric component of the land demand equations (the α's). As expected, household labor force affects land demand positively. An additional adult worker in the household would induce an additional demand for a tenth of a hectare for cultivation. The estimates are statistically the same for both years. The ratio of consumers to workers in the household is also an important determinant of demand for agriculture land. A one-percentage point increase in this ratio results in about 0.005 hectare increase in land demand in both years. This indicates perhaps that households with higher consumption needs per household worker are more likely to choose autoconsomation strategies for income generation. Finally, while the maximum level of education and the age of the household's head have no statistically significant effect on land demand, it appears that, when other things are held constant (e.g., household labor force and dependency ratio), female headed households tend to demand less land than male headed ones. This is perhaps an indication of the lower access to credit and other services faced by woman headed households in rural Mexico.

In addition to computing the non-parametric estimator of the conditional mean $m_t(T_{it})$, we are also interested in estimating changes in the slope $\partial m_t(T_{it})/\partial T_{it}$ at different points in the land endowment space. Since changes in $\partial m_t(T_{it})/\partial T_{it}$ correspond to changes in separability between A_{it} and T_{it}, this gives us a test of the increased separability hypothesis put forward in Section 3. To estimate $\partial m_t(T_{it})/\partial T_{it}$, we employ a smoothed version of the local slope estimator presented in Blundell and Duncan (1998). That is, we first compute the local slope estimator $\partial \tilde{m}_t(T_{it})/\partial T_{it}$, as described in Blundell and Duncan, and then compute its smoothed version by regressing $\partial \tilde{m}_t(T_{it})/\partial T_{it}$ on T_{it}, employing the same Epanechnikov kernel used in the estimation of $m_t(T_{it})$. This procedure was applied to the original sample as well as to the 400 bootstrap samples drawn for the construction of confidence regions around the expected demand estimates. This allowed for the construction of 95% confidence regions around the estimates of $\partial m_t(T_{it})/\partial T_{it}$. The results are presented in Panel B of Figure 3.

As it can be seen, the estimated slopes support the increased separability hypothesis. First, in both 1994 and 1997, the slope for land poor farmers, i.e., those endowed with less than 3

[10] 90% confidence bands were constructed via smooth conditional moments bootstrap (SCM) methods that do not fail under heteroskedasticity (see Gozalo, 1997). 400 pseudo- normal random variables were drawn for each observation to construct the confidence bands.

hectares, is substantially smaller than the slope for land-rich farmers likely to rent land out. The difference between the slopes for farmers with less then 3 and farmers with more than 4 hectares is statistically different from zero at the 5% significance level in both years.[11] This supports the hypothesis that rental market restrictions affected the behavior of landlords more than the behavior of tenants, and therefore, the latter are more likely to be affected by higher rental rates.

In summary, results from the semi-parametric unconditional regression of land used on land owned points to no statistically significant changes in access to land by small owners as well as land supplied by large owners. The results do, however, support the hypothesis of an increased separability between owned and cultivated land by larger farmers. It is nevertheless unclear whether this increased separability results from the removal of land rental regulations, or if changes in the relationship between land endowments and access to credit are the driving force behind this shift. We address this question in the next section.

4.2 A conditional empirical model: Controlling for access to credit

While the impact of reforms depicted in Figure 3 could be caused by the elimination of land confiscation, it could be also brought about by differential changes in access to liquidity. To isolate the impact of lifting rental prohibitions, we derive a test for the *ceteris paribus* impact of reforms for liquidity-constrained households that controls for their access to credit.

A key issue faced by researches interested identifying changes in access to credit is to decide whether or not the observed levels of credit use are supply (access) or demand determined (Kochar, 1991; Conning, 199?; etc.). Fortunately, the survey instruments utilized both in 1994 and 1997 permit us to identify households with high probabilities of being credit contrained. For these households, we can safely assume that the observed level of credit used is a measurement of their individual access to credit S_{it}. Thus, for a subset of households in the sample, we can estimate (4) controlling for access to credit S_{it}.[12]

[11] The SCM bootstrap 95% confidence bounds were not drawn around both curves for the sake of intelligibility.

[12] Households with zero loans were considered credit constrained either if their loan application had been rejected, or if they did not apply knowing that their application would be rejected. Households that receive loans were considered constrained when they declared that they would not be able to receive more credit (or additional loans) at the same terms for the loans received. In total, 153 households were classified as credit constrained in 1994 and 291 in 1997.

Estimators of (4) which utilize only a subset of households are likely to be biased and inconsistent because of endogenous sample selectivity. To see this, consider the following model:[13]

(6.a) $$\begin{aligned} A_{it} &= d_{it} \cdot A^*_{it} = d_{it} \cdot \left[g_t\left(T^*_{it}\right) + \beta S^*_{it} + \gamma' Z^*_{it} + \mu^*_i + \varepsilon^*_{it}\right] \\ &= g_t\left(T_{it}\right) + \beta S_{it} + \gamma' Z_{it} + \mu_{it} + \varepsilon_{it} \end{aligned}$$

(6.b) $$D_{it} = W^D_{it}\theta^D + \omega^D_i + \nu^D_{it}$$

(6.c) $$S_{it} = W^S_{it}\theta^S + \omega^S_i + \nu^S_{it}$$

(6.d) $$d_{it} = 1\{D_{it} - S_{it} \geq 0\}.$$

Here A^*_{it} is the latent land demand of household *i* at time *t*, which is only observed by the econometrician when the household is credit constrained. W^D_{it} and W^S_{it} are vectors of explanatory variables that determine household *i*'s demand for, and access to, credit at time *t* (denoted D_{it} and S_{it}, respectively). ω^D_i, ω^S_i and μ_i are unobservable time-invariant individual-specific effects (possibly correlated with the regressors). ν^D_{it}, ν^S_{it} and ε^*_{it} are unobserved disturbances (not necessarily independent of each other), and θ^S, θ^D, β, and γ are constant parameters. In this model, it is assumed that $\left(d_{it}, W^D_{it}, W^S_{it}, T^*_{it}, Z^*_{it}\right)$ is always observed, while $\left(A^*_{it}, S^*_{it}\right)$ is observed only when $d_{it} = 1$.

Since we are not interested in the parameters θ^S, θ^D per se, we can rewrite model (6) as:

(7.a) $$A_{it} = g_t\left(T_{it}\right) + \beta S_{it} + \gamma' Z_{it} + \mu_{it} + \varepsilon_{it}$$

(7.b) $$d_{it} = 1\{X_{it}\varphi + W_{it}\theta + \omega_i - \nu_{it} \geq 0\},$$

where $X_{it} \equiv \left(T_{it} \vdots Z_{it}\right)$, and W_{it} is a vector of explanatory variables that determine credit supply and/or demand, but do not affect the latent demand for land A^*_{it}. As discussed in Kyriazidou

[13] Note that here we are implicitly assuming that the regression function *g*(.) is such that $d_{it} \cdot g_t\left(T^*_{it}\right) = g_t\left(d_{it} \cdot T^*_{it}\right)$.

(1997), in this set up, it is possible to consistently estimate φ and θ in equation (7.b) using either the conditional maximum likelihood approach proposed by Rasch (1960) and Andersen (1970), or the conditional maximum score method proposed by Manski (1987). However, estimation of $g_t(.)$, β and γ in the main equation of interest (7.a) is confronted with two problems: first, the presence of the unobservable effect $\mu_{it} = d_{it} \cdot \mu_i^*$, and second and more crucial, the potential endogeneity of the regressors $T_{it} = d_{it} \cdot T_{it}^*$ and $Z_{it} = d_{it} \cdot Z_{it}^*$ which arises from their dependence on the selection variable d_{it}, and which may result in "selection bias."

Note that for those observations that are classified as credit constrained in both years (i.e. $d_{i94} = d_{i97} = 1$), the first problem is easily solved by time differencing. First diffecrencing eliminates the effect μ_{it} from equation (5.5.a). Still, application of conventional regression methods to the first-differenced subsample yields inconsistent estimates of the regression function $g_t(.)$ and the parameters β and γ, due to sample selectivity. That is, given the first-differenced equation for households for which $d_{i94}=d_{i97}=1$,

$$(8) \qquad A_{i97}^* - A_{i94}^* = g_{97}\left(T_{i97}^*\right) - g_{94}\left(T_{i94}^*\right) + \beta\left(S_{i97}^* - S_{i94}^*\right) + \gamma'\left(Z_{i97}^* - Z_{i94}^*\right) + \left(\varepsilon_{i97}^* - \varepsilon_{i94}^*\right),$$

and defining the vector $\zeta_i \equiv \left(W_{i94}, W_{i97}, T_{i94}^*, T_{i97}^*, Z_{i94}^*, Z_{i97}^*, S_{i94}^*, S_{i97}^*, \mu_i^*, \omega_i\right)$, there is no reason to expect that $E\left[\varepsilon_{i94}^* \middle| d_{i94} = 1, d_{i97} = 1, \zeta_i\right] = E\left[\varepsilon_{i97}^* \middle| d_{i94} = 1, d_{i97} = 1, \zeta_i\right] = 0$, or that $E\left[\varepsilon_{i94}^* \middle| d_{i94} = 1, d_{i97} = 1, \zeta_i\right] = E\left[\varepsilon_{i97}^* \middle| d_{i94} = 1, d_{i97} = 1, \zeta_i\right]$. Moreover, for each t, the sample selection effect, defined as

$$(9) \qquad \lambda_{it} \equiv E\left[\varepsilon_{it}^* \middle| d_{i94} = 1, d_{i97} = 1, \zeta_i\right] = \Lambda\left(X_{i94}\varphi + W_{i94}\theta + \omega_i, X_{i97}\varphi + W_{i97}\theta + \omega_i, \zeta_i\right),$$

depends not only on the conditioning vector ζ_i, but also on the joint conditional distribution of $\left(\varepsilon_{it}^*, \nu_{i94}, \nu_{i97}\right)$.

To introduce the method developed by Kyriazidou (1997), it is perhaps convenient to assume that $g_t(.)$ is linear and rewrite (6.a) as[14]

$$(10) \qquad A_{it} = \delta_t T_{it} + \beta S_{it} + \gamma' Z_{it} + \mu_{it} + \lambda_{it} + \vartheta_{it},$$

where $\vartheta_{it} \equiv \varepsilon_{it} - \lambda_{it}$ is a new error term, which by construction satisfies $E[\vartheta_{it} | d_{i94} = 1, d_{i97} = 1, \zeta_i] = 0$, and δ_t is a parameter that may or may not vary with t. The basic idea behind Kyriazidou's estimator is to "difference out" the disturbance terms μ_{it} and λ_{it} from (5.8) above. Accordingly, under some weak distributional assumptions regarding $(\varepsilon^*_{it}, \nu_{i94}, \nu_{i97})$ (i.e., *conditional exchangeability*), it can be shown that $(X_{i94}\varphi + W_{i94}\theta - X_{i97}\varphi + W_{i97}\theta) = 0$ implies that $(\lambda_{i94} - \lambda_{i97}) = 0$. Thus, it can be shown that under some regularity conditions (and knowledge of φ and θ), the estimator obtained via OLS computed with a subsample that only contains households for which $(X_{i94}\varphi + W_{i94}\theta - X_{i97}\varphi + W_{i97}\theta) = 0$ and $d_{i94} = d_{i97} = 1$, will be consistent and root-n asymptotically normal. Therefore, Kyriazidou proposes the following two-step estimation procedure that we employ here: In the first step, φ and θ are consistently estimated based on equation (7.b) alone. In the second step, estimates of φ and θ, denoted $\hat{\varphi}$ and $\hat{\theta}$ respectively, are used to estimate the parameters δ_t, β and γ, based on those pairs of observation for which the difference $(X_{i94}\hat{\varphi} + W_{i94}\hat{\theta} - X_{i97}\hat{\varphi} + W_{i97}\hat{\theta})$ is "close" to zero. Specifically, the estimator employed here is given by:

$$(11) \qquad \hat{\Pi} = \left[\sum_{i=1}^{N} \hat{\psi}_i \Delta Y_i' \Delta Y_i \Phi_i\right]^{-1} \left[\sum_{i=1}^{N} \hat{\psi}_i \Delta Y_i' \Delta A_i \Phi_i\right]$$

where $\hat{\Pi} \equiv (\hat{\delta}_t, \hat{\beta}, \hat{\gamma})$, $\Delta Y_i \equiv (\Delta T_i, \Delta S_i, \Delta Z_i)$, $\Phi_i \equiv d_{i1} d_{i2}$, and ψ_i is a weight that declines to zero as the magnitude of the difference $|X_{i94}\hat{\varphi} + W_{i94}\hat{\theta} - X_{i97}\hat{\varphi} + W_{i97}\hat{\theta}|$ increases. As suggested by Kyriazidou, we choose kernel weights of the form:

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[14] In the actual estimation we will assume that $g_t(.)$ is quadratic.

(12) $$\hat{\psi}_i \equiv \frac{1}{h} K\left(\frac{\Delta X_i \hat{\varphi} + \Delta W_i \hat{\theta}}{h}\right)$$

where K is a kernel density function, and h is a bandwidth that shrinks as the sample size increases. As discussed in Kyriazidou, a exclusion restriction is required for the identification of $\Pi \equiv (\delta_t, \beta, \gamma)$. That is, W_{it} in (7.b) must contain at least one element that is not in Z_{it}. In the following section, we describe the estimation strategy and discuss the results.

Estimation strategy and results

In addition to exclusion restrictions, consistent estimates of φ and θ are required for the identification of Π. We compute the conditional fixed-effects logit estimator of φ and θ which is consistent and asymptotically normal under the assumptions that the errors in the selection equation (7.b) are white noise with a logistic distribution and independent of the regressors and the individual effects. As discussed in Kyriazidou, consistency of this conditional fixed-effects logit estimator is necessary for the consistency and asymptotically normality of $\hat{\Pi}$.[15] To satisfy the exclusion restriction we assume that, controlling for access to credit, a household's participation in local formal and/or informal organizations does not affect its demand for land, but do affect its probability of being credit constrained. Therefore, in the selection equation that contains the determinants of being credit rationed, in addition to land endowment and household labor force, we include dummy variables that take the value one if the household is a member of a given organization and zero otherwise. Accordingly, these dummies are not included in the in the land demand equation.[16]

Table 7 below presents the results of the conditional fixed-effects logit estimation of the parameters in (7.b). The dichotomous dependent variable for this regression takes the value one if a household is classified as credit constrained in both 1994 and 1997, and zero otherwise.[17] As it can be seen, the only coefficients that are statistically different from zero at conventional levels

[15] Kyriazidou also suggests a smoothed version of Manski's (1987) conditional maximum score estimator which requires weaker distributional assumptions. For the sake of computational simplicity, however, we opt for the conditional fixed-effects logit estimator.

[16] Describe the different types of organizations...

[17] There were 42 households out of 1286 that reported being credit constrained in both 1994 and 1997.

of significance are the coefficients for land endowments in 1994, the coefficient for the interaction between land endowments and the 1997 dummy variable, and the coefficient for the interaction between the dummy for participation in informal organizations and the dummy for 1997.

The signs and magnitudes of the estimated coefficients for land endowments in 1994 and 1997 are consistent with the conceptual model developed in Section 3. Note that these coefficients measure the impact of land endowments on the probability that a household's demand for credit D_{it} is greater than its access to credit S_{it} (i.e., the impact on $Pr[D_{it}>S_{it}]$), rather than just the impact on access to credit. Thus, right after reforms, larger holders were still cautious about renting land out and therefore demanded more credit than small holders, *ceteris paribus*, since they needed to cultivate more than the "free market optimum" to reduce the risk of land confiscation. On the other hand, before reforms, land could not be used as collateral, and therefore access to credit S_{it} depended very little on land endowments T_{it}. Not surprisingly, our estimates suggest that in 1994 $Pr[D_{it}>S_{it}]$ increased with T_{it}, since D_{it} increased with T_{it}, but S_{it} did not.

With the increased functioning of land rental markets and little or no development in rural credit markets, neither D_{it} nor S_{it} depended much on land endowments T_{it} in 1997. That is, larger holders faced no risky of confiscation and therefore could supply their excess land to the rental market, which implies that credit demand D_{it} did not depend as much on land endowments T_{it} as before. Moreover, with the negligible development of rural credit markets in Mexico, access to credit S_{it} remained independent of land endowments T_{it}. Therefore, as indicated by our estimated coefficients, $Pr[D_{it}>S_{it}]$ did not depend on endowments T_{it} in 1997.[18] Moreover, the positive sign of the estimated coefficient for the 1997 dummy, combined with the practically zero estimated coefficient for the impact of land endowments, suggest that all farmers, regardless of their land endowments, were more likely to be credit constrained in 1997.

Also of interest is the negative coefficient for the participation in informal organizations dummy (-0.749), which is statistically different from zero at the 10% significance level. This

[18] Note that the estimated coefficients for land endowment in 1997 is given by (0.0474 – 0.0461=0.0013) which is not significantly different from zero at the conventional levels.

suggests that informal local organizations do have an impact in helping households cope with liquidity constraint problems. Note that since this estimate is obtained from a first-differenced equation, it is unlikely to be biased due to spurious correlation between participation in such organizations and household-specific time-invariant effects that are related to $Pr[D_{it}>S_{it}]$. Ongoing related work will try to identify whether participation in such informal organizations affects $Pr[D_{it}>S_{it}]$ through its effect on D_{it}, or its effect on S_{it}.

We now discuss the computation of the estimator (11), which utilizes weights computed via the kernel (12) and the first step estimates of φ and θ presented in Table 7. As in Kyriazidou (1997), we utilize a Gaussian kernel for (12). The procedure for computing the optimal bandwidth is thoroughly explained there, and therefore will not be reproduced here. This procedure requires an initial guess h_0, and therefore we compute the estimator $\hat{\Pi} \equiv \left(\hat{\delta}_t, \hat{\beta}, \hat{\gamma}\right)$ for different values of h_0.[19]

Table 8 presents the results for the various initial values h_0. As shown, while the signs of the estimated coefficients remain the same for different values of h_0, their magnitudes vary considerably. Nevertheless, holding all other explanatory variables constant at their median values, the shape of the curve mapping land endowments T onto estimated expected land demand A does not change substantially across different values of h_0.

Panel A in Figure 4 plots this estimated relationship between expected land demand and land endowments. As can be seen, the estimated 1997 curve is considerably flatter than the estimated 1994 curve. This suggests that in 1997 land demand was considerably more equalized across households with similar characteristics (including household labor force and access to credit), but endowed with different amounts of land. That is, between 1994 and 1997, relatively land rich households increased their supply of land to the rental markets, while relatively land poor households increased their demand. As shown in Panel B, the increase in demand by land-poor households and the increase in supply by land-rich households are statistically different

[19] Selection of the optimal bandwidth also requires an assumption regarding the degree of differentiability of the density function of the index $\left(\Delta X_i\hat{\varphi} + \Delta W_i\hat{\theta}\right)$, denoted r, and a constant δ such that $0<\delta<1$. We perform a sensitivity analysis by computing estimates of Π with several values of h_0, r, and δ, and learned that, for a given value of h_0, the estimates changed very little (qualitatively) with different values of r and δ. Therefore, we present the results obtained with r=1, δ=0.5, and h_0=0.5, 1, 2, and 3. For values of h_0 lower than 0.5 and greater than 3, the estimates changed very little.

from zero at the 10% level for households endowed with less than 10 hectares, and for households endowed with more than 30 hectares. Finally, Panel C plots the estimated changes in the slope of land demand *A* with respect to *T*. As shown, this slope appears to have been increased (decreased) for households endowed with less (more) than seven hectares. These changes, however, do not appear to be statistically significant at the 10% level for households with less than 13 hectares. For households with more than 13 hectares, the estimates point to an increased separability between *A* and *T*, as predicted in Section 3.

In summary, once we control for access to credit, the data suggest that land market liberalization reforms have indeed promoted and increased factor-price equalization across households with distinct land endowments via a more active land rental market. This efficiency enhancing effect of better working land rental markets is supported by the statistically significant increase in separability for households endowed with more than 13 hectares. Comparing the results of the unconditional and the conditional analysis, we conclude that decreased access to credit has lessened the efficiency effects of land market reforms. Hence, a clear policy implication is that governments should pursue policies to enhance and improve rural credit markets.

5. Conclusion

This study of the Mexican ejido reforms provides one of the few documented examples where market-friendly reforms appear to have, at the same time, benefited the poor. This is an important indication that the two goals need not be in conflict with each other. Even though our analysis demonstrates that the impact of land market reforms was affected by simultaneous reductions in credit access, this merely serves to demonstrate the importance of proper planning and sequencing of such reforms, rather than –as was feared by critics of these reforms- invalidating the case for measures to improve functioning of rural factor markets.

The finding that, even in an environment where illegal land rental is reported to have been widespread, lifting restrictions on land rental transactions had a positive impact on the poor has important implications for land policy. It suggests that in countries, such as India or the Philippines, where land rental restrictions appear to be more strictly enforced than in Mexico before the reforms, substantial benefits for the poor could be achieved by lifting these restrictions and allowing freer functioning of land rental markets. Identifying ways in which such reforms could be sequenced and implemented without jeopardizing credit access is a challenge for policy as well as research that appears to be well worth taking.

Appendix I

In this appendix we derive the hypotheses regarding the impact of rental market liberalization on the relationship between area cultivated A and land ownership T, which are depicted in Figures 1 and 2 of Section 3. In sum, we are interested on the impact of reforms on the derivative $\frac{\partial A^*}{\partial T}$, where A^*, is the household specific optimal choice of farmed area, and T is the household's land endowment. As indicated in Section 2, under certain regularity conditions, land market liberalization reforms should reduce the difference between this derivative for households participating the two possible land market participation regimes: (i) *renting-in* or *tenant-farmer* regime (for which $A>T$), and (ii) *renting-out* or *landlord* regime (for which $A<T$). That is, $\frac{\partial A^*_{A<T}}{\partial T}$ should become closer to $\frac{\partial A^*_{A>T}}{\partial T}$, for liquidity constrained and unconstrained households.

As stated, the household's decision problem from which we derive this relationship is given by:

Maximize: $c + W(A,T;M)$

w.r.t: X_f, X_h, A, and X_o.

s.t.:

$$c = F(E,A) - rA - wX_h + rT + wX_o$$

$$\overline{L} \geq X_f + X_o$$

$$E = E(X_f, X_h)$$

$$W(A,T;M) = \begin{cases} \overline{\rho}\omega\left(\frac{A}{T}\right)V(M-T) + \left[1-\overline{\rho}\omega\left(\frac{A}{T}\right)\right]V(M), & \textit{for} \quad A<T \\ V(M) & \textit{otherwise} \end{cases}$$

$$rA + wX_h \leq S + wX_o + rT$$

$$X_f, X_h, A, X_o \geq 0 \qquad ,$$

where $F(.)$ is a constant returns to scale production (CRTS) function (and therefore is homogeneous of degree 1), and the effort function $E(.)$ is given by:

(1) $$E = \left(X_f + X_h\right)\left[\frac{X_f}{X_f + X_h}\right]^{1-\gamma}$$

which is also homogeneous of degree one. Before reforms, for households operating in the rent-out (landlord) regime, the utility for terminal wealth $W(A,T;M)$ is such that:

(2) $$\frac{\partial W}{\partial A} \equiv W_A \equiv -\left(\frac{V(M)-V(M-T)}{T}\right)\overline{\rho}\omega' > 0\,,$$

and:

(3) $$\frac{\partial W}{\partial A \partial A} \equiv W_{AA} \equiv -\left(\frac{V(M)-V(M-T)}{T^2}\right)\overline{\rho}\omega'' < 0$$

(4) $$\frac{\partial W}{\partial A \partial T} \equiv W_{AT} \equiv -\left[\frac{V_T T - \left(V(M)-V(M-T)\right)}{T^2}\right]\overline{\rho}\omega' + \left[\frac{V(M)-V(M-T)}{T^2}\right]\frac{A}{T}\overline{\rho}\omega'' > 0\,.$$

The signs from (2), (3) and (4) follow from the fact that $V(.)$ is monotonically increasing and concave, which implies that $[V(M) - V(M-T)]>0$, and that $\omega'<0$, and $\omega''>0$, by assumption.

As seen in the maximization problem above, in addition to the two land market participation regimes, depending on whether the liquidity constraint $rA + wX_h \leq S + wX_o + rT$ binds or not, households may also operate in a liquidity-unconstrained (when $rA + wX_h < S + wX_o + rT$), or in a liquidity-constrained regime (when $rA + wX_h = S + wX_o + rT$). For each of these four regimes, we denote the optimal choice of area cultivated by:

$A^* = A^u_{j,A>T}\left(T; w, r, \overline{L}\right)$, for liquidity-unconstrained tenant-farmers ,

$A^* = A^u_{j,A<T}\left(T; w, r, \overline{L}\right)$, for liquidity-unconstrained landlord farmers,

$A^* = A^c_{j,A>T}\left(T; w, r, \overline{L}\right)$, for liquidity-constrained tenant-farmers , and,

$A^* = A^c_{j,A<T}\left(T; w, r, \overline{L}\right)$, for liquidity-constrained landlord farmers.

The subscript *j* indicates whether the household the choice A^* is made before reforms (*j*=0), or after reforms (*j*=1). We start examining the case of liquidity-unconstrained households.

Liquidity-unconstrained regime

We start by noting that, because *F*(.) exhibits CRTS, a liquidity-unconstrained household that faces an identical wage rate *w* for both buying and selling labor, would <u>not</u> be observed concurrently cultivating and selling labor to off-farm activities. That is, without limits to the amount of liquidity and land that can be rented at market interest and rental rates, households would either sell all their labor endowments to the market (in which case $X_o = \overline{L}$ and *A*=0), or they would instead rent in enough land so that $X_f = \overline{L}$, and $X_o = 0$. Therefore, since our sample contains only cultivating households for which *A*>0, we consider only the case of unconstrained households for which $X_o = 0$, and $X_h > 0$, i.e., liquidity-unconstrained labor-buyers. In this case, the relevant first order necessary conditions (FONCs) for the optimality of A^* before reforms are given by:

(5)
$$\begin{aligned} &F_E E_h - w = 0 \\ &F_A - r + W_A = 0 \end{aligned}.$$

As it can be shown, because of the strict-concavity of *F*(.), and the assumed functional forms of *E*(.) and *W*(.), the Hessian matrix of second derivatives is negative-definite, and therefore, (5) are not only necessary but are also sufficient conditions for a maximum. Hence, by Cramer's rule, the derivative of A^* with respect to land owned *T* is positive and given by:

$$\frac{\partial A^*}{\partial T} = \frac{\begin{vmatrix} F_{EE}E_h^2 + F_E E_{hh} & 0 \\ 0 & -W_{AT} \end{vmatrix}}{|H|} = -\frac{(F_{EE}E_h^2 + F_E E_{hh})W_{AT}}{(F_{EE}E_h^2 + F_E E_{hh})W_{AA} + F_E E_{hh} F_{AA}} \geq 0$$

Since $W_A = W_{AA} = W_{AT} = 0$ for households operating in the rent-in regime before reforms, and for all households after reforms, we conclude that:

$$\frac{\partial A^u_{0,A>T}}{\partial T} = \frac{\partial A^u_{1,A>T}}{\partial T} = \frac{\partial A^u_{1,A<T}}{\partial T} = 0 < \frac{\partial A^u_{0,A<T}}{\partial T},$$

which explains the reform induced shifts in A^* depicted in Figure (??).

Liquidity constrained regime

In contrast to households operating in the credit unconstrained regime, households that face a binding liquidity constraint may be observed concurrently cultivating positive areas of land (A>0) and selling labor to the market (X_o>0). Therefore, in addition to the two land market participation regimes ($A>T$ and $A<T$), cultivating households that are liquidity-constrained can be observed in on of two labor market participation regimes: (i) labor sellers (X_o>0, X_h=0), and (ii) labor buyers (X_o=0, X_h>0). We start the analyzes with labor sellers.

Liquidity-constrained labor sellers:

For liquidity-constrained labor sellers, the relevant FONCs are:

$$F_E - (1+\lambda)w = 0$$
$$F_A - (1+\lambda)r + W_A = 0$$
$$S + w(\overline{L} - X_f) + rT - rA = 0$$

As it can be shown, because of the assumptions regarding $F()$, $E(.)$ and $W(.)$, the FONCs are also sufficient conditions for a constrained maximum. The determinant of the bordered-Hessian is positive and given by:

$$\left|\overline{H}\right| = \frac{F_{AE}}{AE}\left(wE + rA\right)^2 - w^2 W_{AA} > 0,$$

since F_{AE}>0, by the homogeneity of degree one of $F(.)$, and because W_{AA}<0 by assumption. Therefore, by Cramer's rule we have:

$$\frac{\partial A^*}{\partial T} = \frac{\frac{F_{AE}}{E} r\left(wE + rA\right) + w^2 W_{AT}}{\frac{F_{AE}}{AE}\left(wE + rA\right)^2 - w^2 W_{AA}} \geq 0, \tag{6}$$

since $W_{AT}>0$. Thus, for all households, before and after reforms, we would have:

(7) $$\frac{\partial A^u_{0,A>T}}{\partial T}, \frac{\partial A^u_{0,A<T}}{\partial T}, \frac{\partial A^u_{1,A>T}}{\partial T}, \frac{\partial A^u_{1,A<T}}{\partial T} > 0 .$$

Note however, that in the neighborhood of $A^*=T$ we have that $|W_{AT}|>|W_{AA}|$, because the first term in (4) is positive, and the second term is equal to $|W_{AA}|$. Therefore, before reforms we have (in the neighborhood of $A^*=T$):

(8) $$\frac{\partial A^u_{0,A<T}}{\partial T} > \frac{\partial A^u_{0,A>T}}{\partial T} > 0 ,$$

which explains the "kink" depicted in Figure (??) around $A^*=T$.

Obviously, because $W_A=W_{AA}=W_{AT}=0$ for all T after reforms, we have:

(9) $$\frac{\partial A^u_{1,A<T}}{\partial T} = \frac{\partial A^u_{1,A>T}}{\partial T} > 0 ,$$

in the neighborhood of $A^*=T$, as depicted in Figure 2.

Liquidity-constrained labor buyers:

For liquidity-constrained labor buyers, the relevant FONCs are:

$$\begin{aligned} &F_E E_h - (1+\lambda)w = 0 \\ &F_A - (1+\lambda)r + W_A = 0 \\ &S + rT - rA - wX_h = 0 \end{aligned} .$$

Again, because of the assumptions regarding $F()$, $E(.)$ and $W(.)$, the FONCs are also sufficient conditions for a constrained maximum, and therefore, by Cramer's rule we have:

$$\frac{\partial A^*}{\partial T} = \frac{\frac{F_{AE}}{E} rE_h\left(wE + rE_h A\right) + w^2 W_{AT} - r^2 F_E E_{hh}}{\frac{F_{AE}}{AE}\left(wE + rE_h A\right)^2 - w^2 W_{AA} - r^2 F_E E_{hh}} \geq 0 .$$

As in the case of liquidity-constrained labor buyers, it is straightforward to show that in the neighborhood of $A^*=T$, $|W_{AT}|>|W_{AA}|$, and therefore, before reforms we have:

(10)
$$\frac{\partial A^u_{0,A<T}}{\partial T} > \frac{\partial A^u_{0,A>T}}{\partial T} > 0 ,$$

in the neighborhood of $A^*=T$, which is consistent with the "kink" depicted in Figure (??). Because $W_A=W_{AA}=W_{AT}=0$ for all T after reforms, here we also have that:

(11)
$$\frac{\partial A^u_{1,A<T}}{\partial T} = \frac{\partial A^u_{1,A>T}}{\partial T} > 0 .$$

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Figure 1: The Impact of Reforms in the Absence of Capital Contstraints

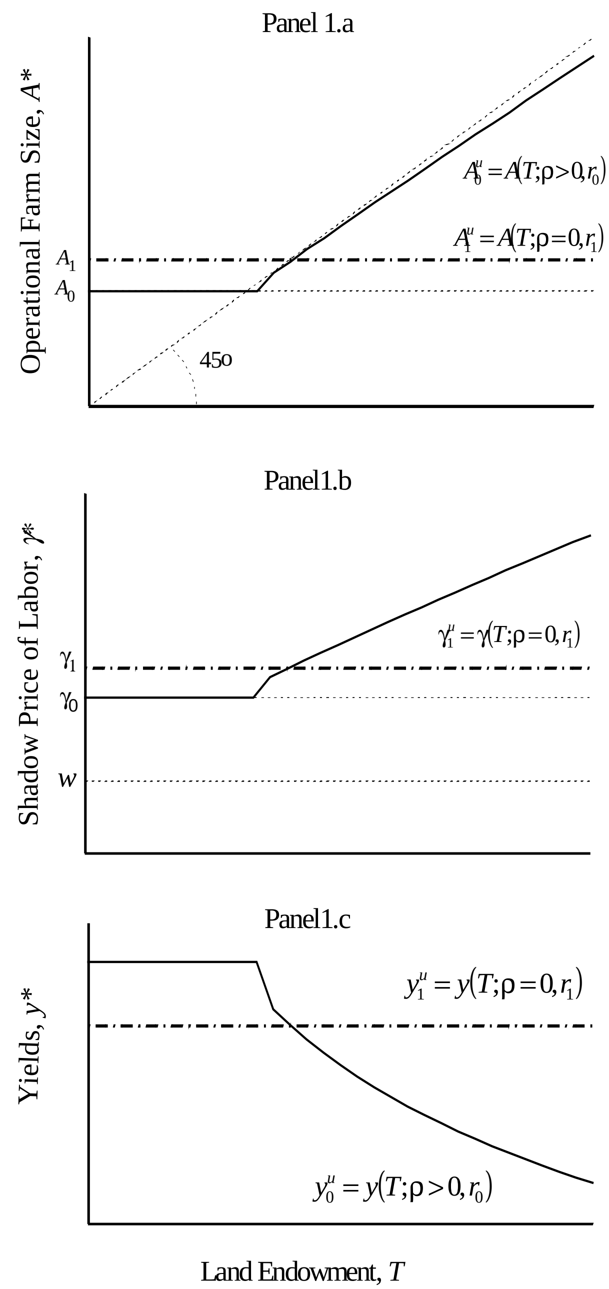


Figure 2: The Impact of Reforms under Capital Contstraints

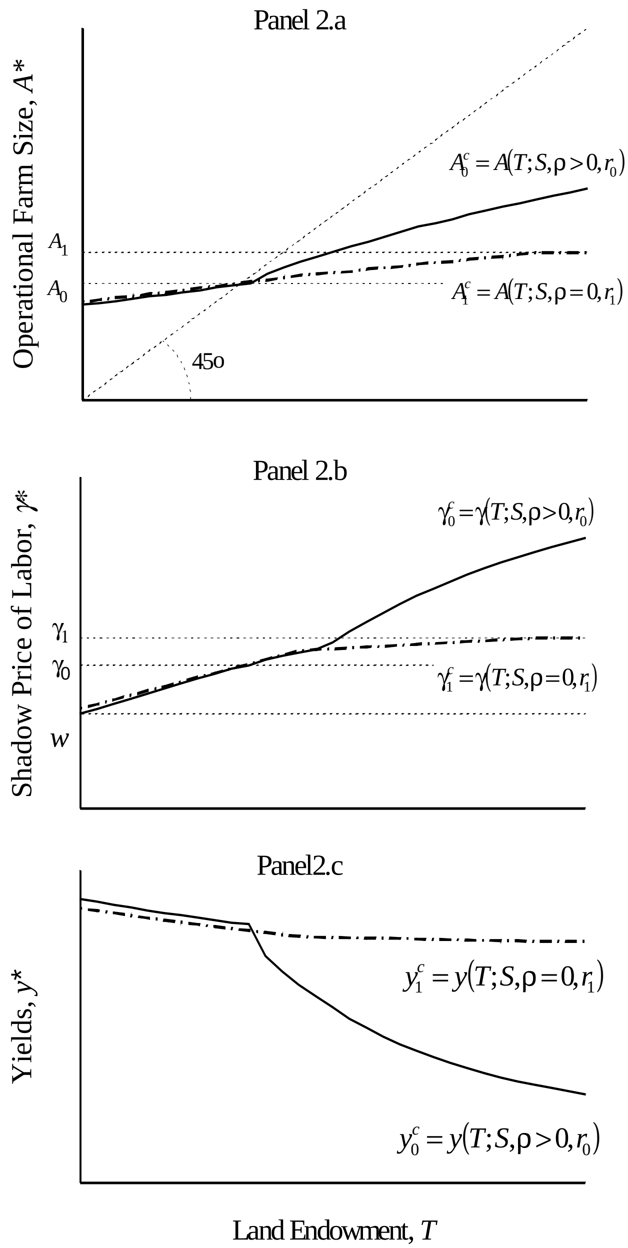
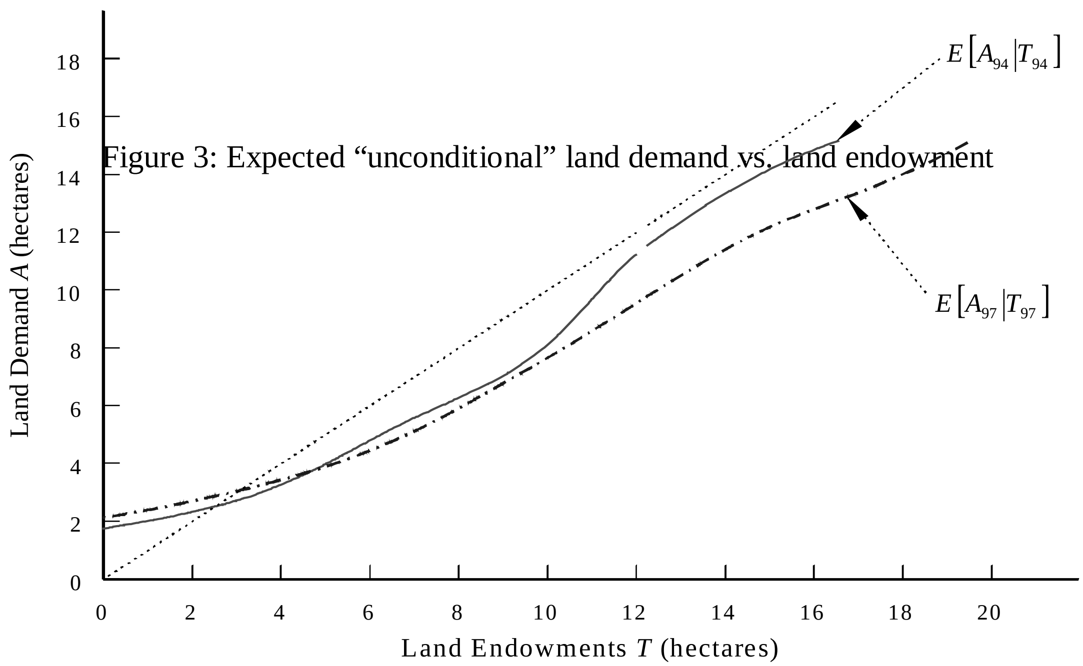


Figure 3: Expected “unconditional” land demand vs. land endowment

Panel A



Panel B

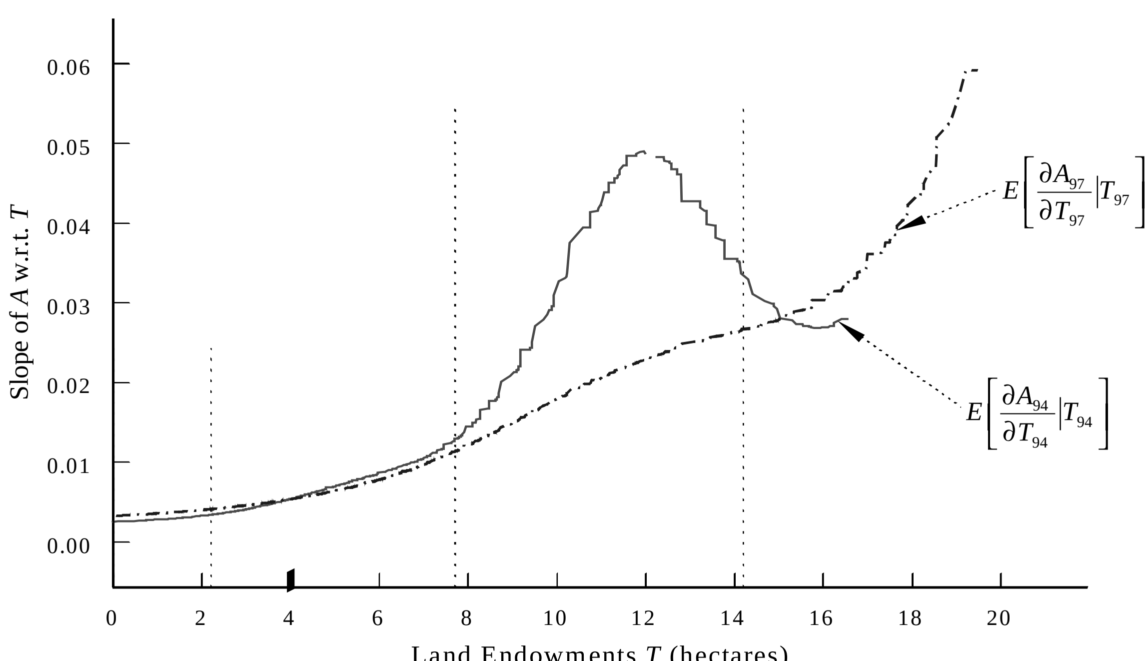
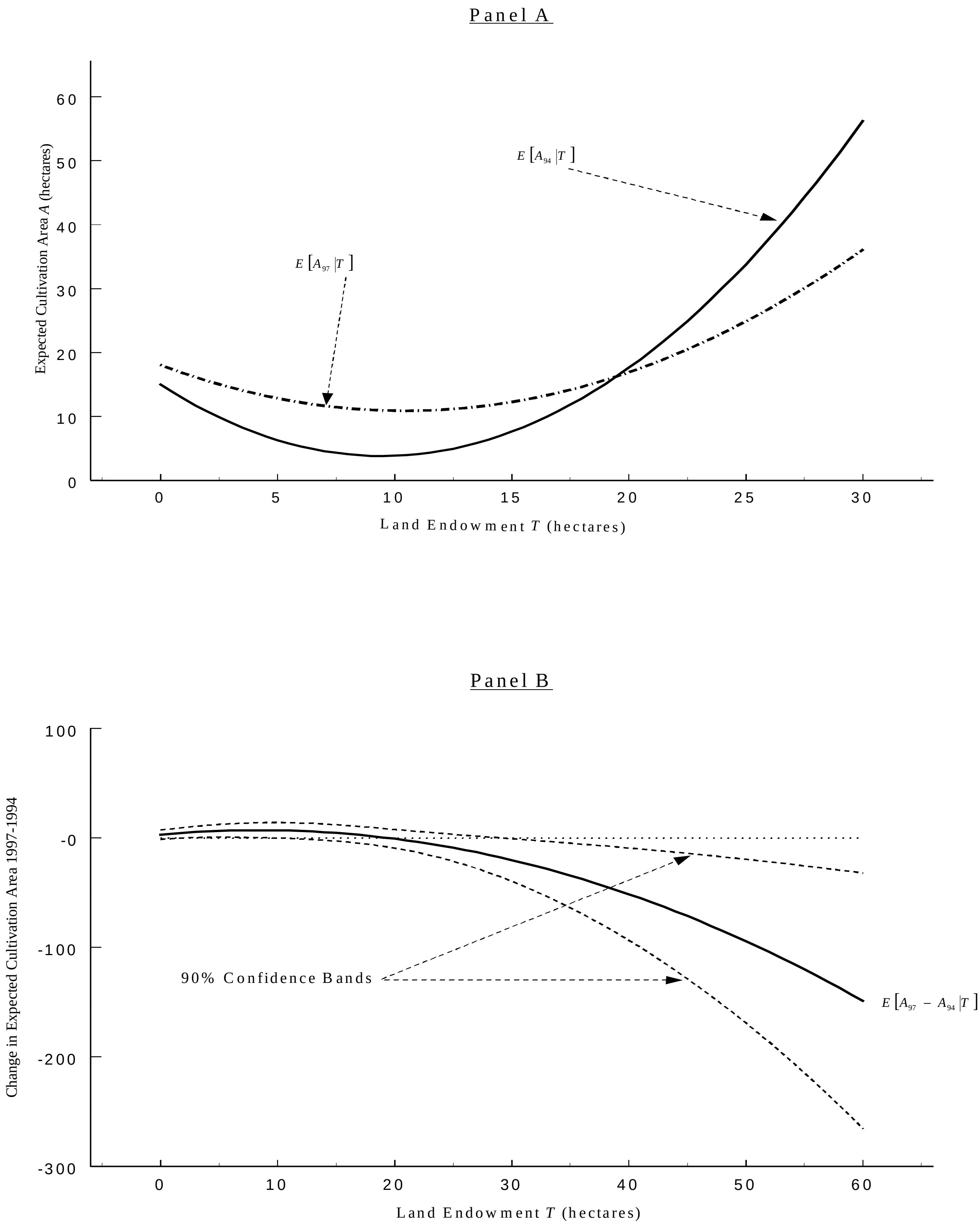
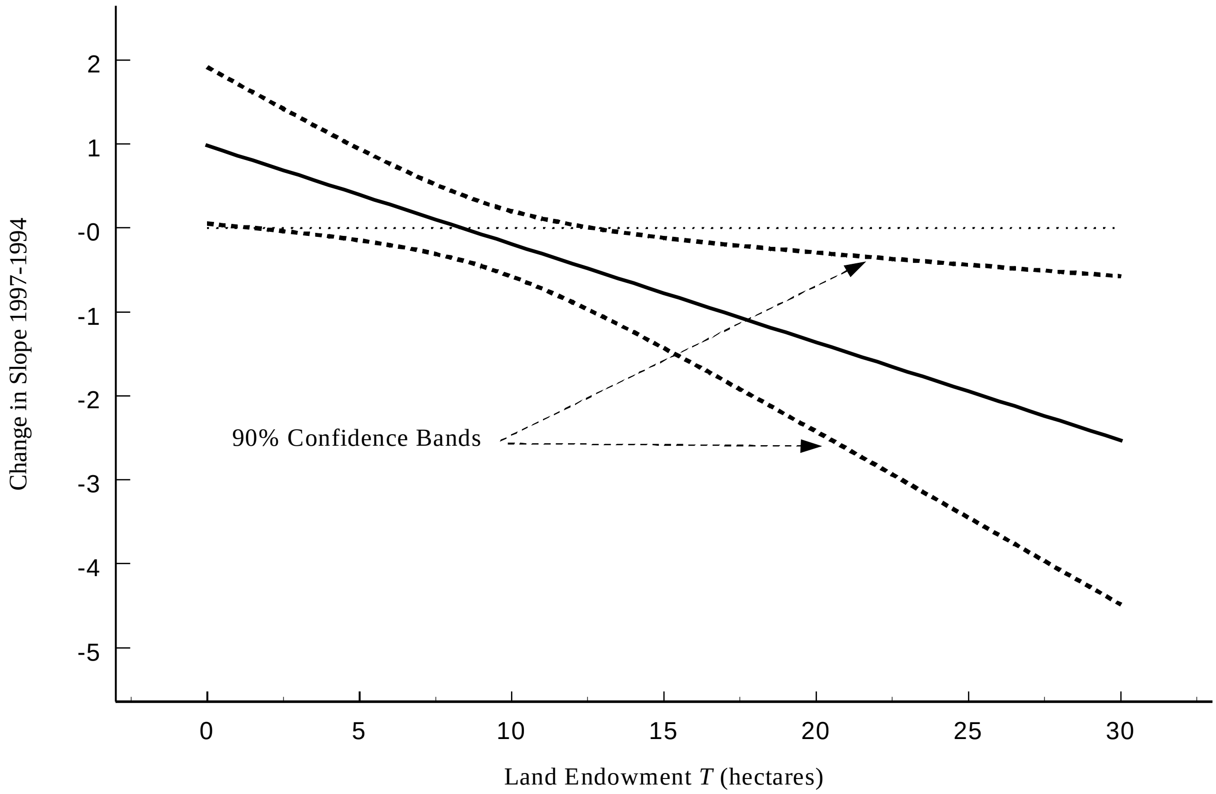


Figure 4: Expected “conditional” land demand vs. land endowments



Panel C



Change in Slope 1997-1994
2
1
-0
-1
-2
-3
-4
-5
90% Confidence Bands
0
5
10
15
20
25
30
Land Endowment T (hectares)

Table 6: Parametric estimates of Land Demand Equation (Dependent variable = Cultivated area in NRE hectares)

Explanatory Variables	1994		1997	
	Estimates	*Std. Errors*	*Estimates*	*Std. Errors*
Household labor force (adult equivalent units)	0.0870*	0.0514	0.1069**	0.0339
Dependency ratio (consumers/workers)	0.5548**	0.0492	0.4278**	0.0201
Maximum education in household (years)	0.0706	0.0821	0.0095	0.0061
Age of Head of Household	-0.0005	-0.2662	0.0010	0.2757
Head's Age Squared	0.0472	0.2252	-0.1226	-0.3062
Dummy for female headed household	-1.7218**	-0.1080	-0.6105**	-0.0192
Dummy for North region (South Pacific region excluded)	-0.3675**	-0.0501	0.6315**	0.0347
Dummy for North Pacific region	-1.6469**	-0.1828	0.4408**	0.0238
Dummy for Center region	-0.3328**	-0.0499	-0.0705**	-0.0047
Dummy for Gulf region	-0.1259**	-0.0158	0.0348**	0.0021

Table 7: Conditional fixed-effects logistic regression of the probability of being credit constrained

EXPLANATORY VARIABLES	Coefficient	Std. Err.
Land Endowment (hectares NRE owned)	.0474**	.0151
Household Labor Force (number of adult equivalent)	.0017	.0677
Dummy for member of informal organization[20]	.1212	.2872
Dummy for member of class organization[21]	.6534	.4853
Dummy for member of single purpose organization[22]	.0757	.3326
Shift between 1994 and 1997 (α_{97} - α_{94}):		
1997 Dummy	.8472**	.3535
Land Endowments ×1997 Dummy	-.0461**	.0136
Household Labor Force	.0758	.0832
Dummy for member of informal organization ×1997 Dummy	-.7487*	.4131
Dummy for member of class organization ×1997 Dummy	-.2894	.6232
Dummy for member of single purpose organization ×1997 Dummy	.0923	.4427
Observations	1286 Hhs × 2 periods	
Log likelihood	-242.10	

** indicates statistically different from zero at the 5% significance level
* indicates statistically different from zero at the 10% significance level

[20] Informal organizations are …

[21] Class organizations are …

[22] Single purpose organizations are…

Table ??: Parameter estimates of land demand equation for credit constrained households

	Results for different constants h_0 used in computing optimal bandwidth							
EXPLANATORY VARIABLES	h_0=.5		h_0=1		h_0=2		h_0=3	
	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.	Coeff.	Std. Err.
Land Owned (hectares NRE)	-0.1396	2.1342	-0.3716	2.1879	-1.3766	2.2055	-0.6878	2.2059
Land Owned^2 (hectares NRE)	0.0365	0.0504	0.0400	0.0530	0.0660	0.0557	0.0474	0.0545
Labor Force (number of adult equival.)	0.7512*	0.3935	0.6555*	0.3882	0.5467	0.3693	0.5971	0.3810
Max. Education in household	-0.1294	0.1791	-0.1760	0.1722	-0.2002	0.1595	-0.1986	0.1668
Age of head of household	-0.6928*	0.4033	-0.6381	0.4044	-0.5302	0.4030	-0.5927	0.4051
Sex of head of household	6.2760	6.6906	6.6874	6.4741	6.3709	6.2825	6.6810	6.3690
Total supply of credit to household	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001	0.0001
Age of head of household squared	0.0069*	0.0037	0.0064*	0.0037	0.0054	0.0038	0.0060	0.0038
Shift between 1994 and 1997 (β_{94}-β_{97}):								
Dummy 1994	-12.168**	4.3828	-7.4875**	3.4600	-3.0658	2.6150	-4.9120*	2.9341
Land Owned × Dummy 1994	-0.5472	0.5466	-0.6802	0.5500	-0.9829*	0.5663	-0.7896	0.5541
Land Owned squared × Dummy 1994	0.0392*	0.0235	0.0429*	0.0252	0.0585**	0.0280	0.0477*	0.0264
Labor Force × Dummy 1994	0.1826	0.7531	0.3114	0.7419	0.3509	0.7344	0.3597	0.7368
Max. Education in HH × Dummy 1994	0.1301	0.2088	0.1762	0.2005	0.1889	0.1817	0.1973	0.1937
Age of Head of HH × Dummy 1994	0.8749**	0.3481	0.7186**	0.3343	0.5642*	0.3195	0.6305*	0.3273
Sex of Head of HH × Dummy 1994	-6.4765	10.7813	-8.3716	10.5094	-9.0426	10.2532	-9.1207	10.3668
Supply of Credit to HH × Dummy 1994	-0.0004**	0.0001	-0.0005**	0.0001	-0.0005**	0.0001	-0.0005**	0.0001
Age of Head of HH^2 × Dummy 1994	-0.0090**	0.0032	-0.0076**	0.0031	-0.0062**	0.0030	-0.0068**	0.0030

** indicates statistically different from zero at the 5% significance level
* indicates statistically different from zero at the 10% significance level